Economic Modelling 29 (2012) 2504–2513

Contents lists available at SciVerse ScienceDirect

Economic Modelling

journal homepage: www.elsevier.com/locate/ecmod

A threshold cointegration analysis of interest rate pass-through to UK mortgage rates

Ralf Becker [a,*], Denise R. Osborn [a], Dilem Yildirim [b]

[a] Centre for Growth and Business Cycle Research, Economics, School of Social Sciences, The University of Manchester, Oxford Rd, Manchester M13 9PL, UK
[b] Economics, Middle East Technical University, Ankara, Turkey

ARTICLE INFO

Article history:
Accepted 4 August 2012

JEL classification:
C51
C52
G21

Keywords:
Interest rate transmission
Mortgage rates
Nonlinear cointegration

ABSTRACT

This paper empirically analyses the interest rate transmission mechanism in the United Kingdom by exploring the pass-through of the official rate to the money market rate and of the market rate to the mortgage rate. Potential asymmetries, due to financial market conditions and monetary policy, lead to the use of a nonlinear threshold error-correction model, with hypothesis tests based on nonstandard bootstrap procedures that take into account the discrete nature of changes in the official rate. The empirical results indicate the presence of substantial asymmetries in both steps of the process, with these asymmetries depending on past changes in the money market rate and whether these are motivated by official rate changes. Generalized impulse response function analysis shows that adjustments differ with regard to the sign and magnitude of interest rate changes in a way that is consistent with conditions in the interbank and mortgage markets over the recent period.

1. Introduction

The principal tool of monetary policy, as conducted by many central banks in developed and developing countries around the world, is the official short-term interest rate. By varying its official rate, the central bank aims to influence the retail loan and deposit rates offered by commercial banks to non-financial institutions and individuals, in order to achieve its aims for inflation and output. However, the "pass-through" from official to commercial interest rates is neither necessarily immediate nor one-to-one. Indeed, it became evident during the recent credit crunch that the money market itself plays a key role in the interest rate transmission process, with the rates at which commercial banks provide short-term loans to each other in this market reflecting demand and supply considerations, as well as the current official interest rate.

Despite the importance of retail rates in determining the effectiveness of monetary policy, there is a surprisingly scant literature on the pass-through from official to retail interest rates.[1] Nevertheless, recent empirical contributions (including Fuertes et al., 2010; Hofmann and Mizen, 2004; Payne, 2007; Sander and Kleimeier, 2004) find strong evidence of nonlinearities, with retail rates responding asymmetrically to disequilibrium in relation to the official rate or its proxy. However, these studies typically ignore the role of the money market.

The historically high spread for money market rates over official rates at certain times during the credit crunch has highlighted the crucial role of these markets for determining both the level of retail interest rates and the availability of funds. While the operation of the money market has undeniably been affected by the abnormal conditions of the recent credit crunch, this has also served to emphasise the lack of research to date about the nature of the pass-through from official interest rates to money market rates and how these, in turn, affect retail rates.

Introducing money market considerations points to a two-stage transmission process, namely from official rates to money market rates and from money market rates to retail interest rates. The only study that considers such a two-stage process is de Bondt (2005), who examines this through a three equation linear system, consequently not allowing for the nonlinearities found in other pass-through studies.

The present paper analyses the interest rate transmission mechanism in the United Kingdom by exploring both the pass-through of the official rate to the money market rate and subsequently the money market rate to the retail mortgage rate in a nonlinear context. The mortgage rate is selected for study since it is the key interest rate in terms of household expenditure and, consequently, is the "headline" rate used by the press for interpreting the impact of monetary policy changes by the Bank of England. Two sample periods are used in our analysis, namely one ending early in 2006 that does not include the credit crunch and may be considered a "normal" period and an extended sample to August 2008 that includes a period in which the credit market was under considerable stress (credit crunch).

Methodologically, a threshold cointegration relationship is employed, in line with other pass-through studies. Unlike previous studies, however,

* Corresponding author. Tel.: +44 161 2754807.
E-mail addresses: ralf.becker@manchester.ac.uk (R. Becker), denise.osborn@manchester.ac.uk (D.R. Osborn), dilem@metu.edu.tr (D. Yildirim).

[1] See de Bondt (2005), who provides a useful summary of the literature relating to individual euro area countries.

0264-9993/$ – see front matter © 2012 Elsevier B.V. All rights reserved.
http://dx.doi.org/10.1016/j.econmod.2012.08.004

both pass-through stages are analyzed in this context. Further, our methodology employs tests for nonlinearity that recognize the inherent unidentified parameter problem, in the spirit of Balke and Fomby (1997), Enders and Siklos (2001) and Hansen and Seo (2002). Indeed, the application of such tests in our context requires the development of a new bootstrap testing procedure, due to the discrete nature of changes in the official Bank of England rate. All previous UK studies ignore this important characteristic of the data and apply tests that assume continuous variables.

The remainder of the paper is organised as follows. Section 2 reviews the background literature, while Section 3 describes our data. The econometric methodology is discussed in Section 4. Substantive empirical results together with generalized impulse response functions (Koop et al., 1996) that facilitate interpretation are presented in Sections 5 and 6, respectively. Section 7 contains some concluding remarks.

2. Previous literature

Money market rates are the marginal costs of funds faced by banks. However, due to adjustment costs (namely the costs to banks of changing mortgage rates), banks may not adjust mortgage rates in response to very small market rate changes and/or changes that are expected to be temporary. Hence, when they have some monopolistic power, banks may wait for large changes and/or a sequence of small changes to accumulate, leading to asymmetry in the pass-through to retail rates. Although discussed in the context of base rate changes, a theoretical model of this type is developed by Hofmann and Mizen (2004). The setting of retail rates is also examined in the context of increasing market deregulation, with competition leading to a more complete and symmetric pass-through by increasing the cost of not adjusting (Corvoisier and Gropp, 2001).

Empirical analysis of asymmetries in the interest rate pass-through dates back to Neumark and Sharpe (1992), who apply a partial adjustment model with differing adjustment speeds depending on the sign of the disequilibrium. Employing nonlinear threshold error-correction models (ECMs), Burgstaller (2005) and de Bondt et al. (2005) examine mortgage rates in Austria and the euro zone, respectively, and find different responses to positive and negative deviations from equilibrium. However, both studies make the untested assumption that the threshold value giving rise to nonlinearity is zero.

Sander and Kleimeier (2004) and Payne (2007) apply the testing methodology of Enders and Siklos (2001) in order to allow for an endogenously determined threshold and find an asymmetric pass-through for variable mortgage rates in the Euro area and United States, respectively. Nevertheless, other work (Payne, 2006a, 2006b) concludes that adjustments for US fixed and 30-year mortgage rates are symmetric.

Although early studies of UK mortgage rates (Heffernan, 1993; Paisley, 1994) assume linearity, more recent analyses find significant asymmetries in the pass-through from the official rate to retail rates; see Heffernan (1997), Hofmann and Mizen (2004) and Fuertes et al. (2010). Heffernan finds that the mortgage rate reacts more slowly when the official rate is rising than when it is falling but, in contrast, using a later sample period and more disaggregated data, Fuertes et al. (2010) find quicker responses to rising official rates. This latter paper also uncovers faster adjustment for larger changes in the official rate, while Hofmann and Mizen (2004) detect faster adjustment when the deviation from equilibrium is widening or expected to widen, implying possible nonlinearities associated with size effects should also be examined.

A common finding of the above studies is that the pass-through from official or money market rates to mortgage rates is incomplete in the long run, and hence mortgage rates do not fully reflect the effects of monetary policy as conducted by the central bank.

Despite these studies, the literature on the dynamics of the pass-through from official rates to money market rates is relatively thin. However, the findings of Kuttner (2001) emphasize the different impacts on money market rates of anticipated versus unanticipated monetary policy actions by the Federal Reserve, while Sarno and Thornton (2003) uncover strong evidence of a nonlinear adjustment between the federal funds rate and the 3 month Treasury bill rate.

This study incorporates both stages of the pass-through for the UK in a threshold ECM framework. This raises methodological issues for hypothesis testing, due to the discrete nature of interest rate changes implemented by the Bank of England.[2] We confront these issues by developing a new simulation-based procedure that explicitly recognises the discrete nature of this variable. Before detailing this methodology, the data are examined in the next section.

3. Data

This study employs interest rate series measured at the end of month from January 1995 to August 2008. The starting point is dictated by changes to the structure of mortgage lending in the UK and the consequent availability of consistent data series on mortgage rates from the Bank of England.[3] Prior to deregulation in the 1980s, the UK mortgage market was dominated by building societies who effectively operated an interest rate cartel. Deregulation brought the large-scale entry of banks into the market, with further legislation passed in the mid-1990s to ensure that building societies were able to compete within a relatively equitable competitive environment; Stephens (2007) provides a detailed discussion and analysis of these changes. The starting date of 1995 also coincides with a period of stability in UK monetary policy since the adoption of inflation targeting in October 1992. Although full independence was given to the Bank of England only in May 1997, researchers interested in the nature and impact of UK monetary policy typically find the period from around 1992 to be a single regime (for example, Benati, 2004; Kesriyeli et al., 2006).

The sample used in this paper ends in August 2008. The placement of Fannie Mae and Freddie Mac in US federal conservatorship and the collapse of Lehman Brothers in September 2008 triggered unprecedented intervention of governments and central banks into the banking sector and hence into money and credit markets. Since standard models cannot capture these extraordinary events, the period after September 2008 is excluded from our analysis.

We are alert that the back end of our sample includes the beginning of the credit crunch and hence some of our results are based on the shorter sample period of January 1995 to January 2006, which we judge to be relatively unstressed. Therefore, whenever appropriate, we also comment on the robustness of results to the extension of this sample period to August 2008 and, when the extended sample results clarify recent developments, we comment more thoroughly on these. Nevertheless, the primary aim of the paper is to investigate the interest rate pass-through mechanism in “normal” times and recent events are used only to provide robustness checks.

Ideally a longer sample period would be employed for an analysis that involves both long-run and nonlinear modelling. However, our baseline sample (to January 2006) contains 133 monthly observations and the extended sample (to August 2008) 164 observations. Further, our estimated models are quite parsimonious, so we are confident that sufficient data are available for our results to be reliable; indeed, we use the extended data to August 2008 to provide some reassurance on this.

[2] The empirical analysis of de Bondt (2005) replaces the official rate with an overnight rate, apparently to avoid the discrete and infrequent changes exhibited by the official series (de Bondt, 2005, p.48).

[3] All data are from the Bank of England database www.bankofengland.co.uk/statistics/index.htm.

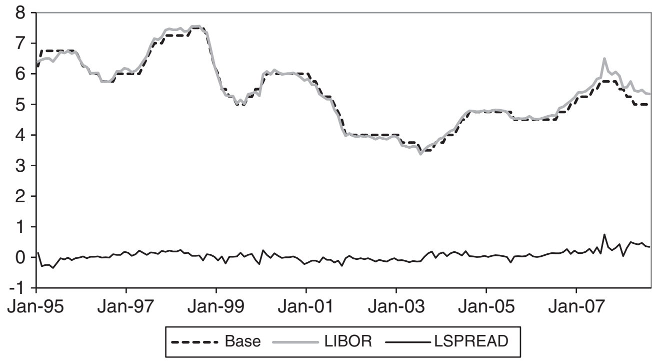


Fig. 1. The base rate and 1 month London inter-bank offer rate (LIBOR), together with the difference between Libor and the base rate (LSPREAD), are shown over the extended sample from January 1995 to August 2008.

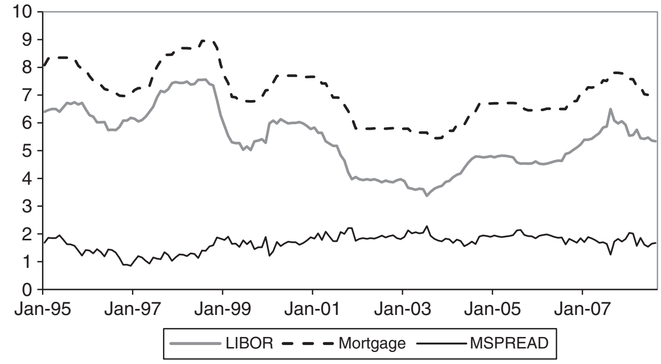


Fig. 2. One month London inter-bank offer rate (LIBOR) and average standard variable mortgage rate of banks (mortgage rate), together with the difference between the mortgage rate and Libor (MSPREAD), are shown over the extended sample period from January 1995 to August 2008.

Fig. 1 shows the official Bank of England interest rate (or base rate),[4] together with the 1 month London inter-bank offer rate (denoted as LIBOR) and the spread between these rates (LSPREAD). Overall, Fig. 1 gives the impression of complete, or near complete, pass-through of official rates to the money market. Only during the final part of this period did the money market rate trade at a persistent positive premium relative to the base rate. The discrete nature of the official rate is also evident from Fig. 1, with monetary policy typically implemented by the Bank through one quarter (or 25 basis points) changes in the rate. Indeed, there are relatively long sequences where the rate is unchanged, notably the 15 months from November 2001. On a relatively small number of occasions, the official rate changes by ± 0.50, but no larger shifts are observed over this period.

This study uses the LIBOR rate as the market rate, since this is the reference rate for (sterling) borrowing and lending in the London interbank market; it is also the market rate employed in the UK studies of Paisley (1994) and Heffernan (1993). Although the LIBOR is calculated for a range of maturities, from overnight to 12 months, the 1-month maturity rate is selected for analysis since this shows the highest correlation (for both levels and changes) with the mortgage rate. Our use of correlation analysis to select the market rate that provides the appropriate marginal cost measure follows de Bondt (2005).

The mortgage rate is the average standard variable mortgage rate (SVR) of banks,[5] which reflects the general rate of interest paid by borrowers. Miles (2004) indicates that at the end of 2003 around 35% of mortgage loans were at standard variable rate, while fixed and discounted variable mortgages made up around 25% and 18% of total loans, respectively. As seen in Fig. 2, LIBOR and the mortgage rate are highly correlated, although with a significant mark-up (MSPREAD) that generally fluctuates between 1 and 2 percentage points. The data seem to suggest that this spread widens in the latter part of the period. Despite the high correlation of 0.971 between the mortgage rate and the LIBOR, the former is a little less volatile with a standard deviation of 0.986 compared with 1.180 for LIBOR. Moreover, the correlation between changes in the mortgage and LIBOR rates is substantially lower, at 0.457, than the correlation between changes in the base rate and LIBOR, which is 0.665. This suggests that the pass-through from LIBOR to the mortgage rate may be imperfect.[6]

A conventional unit root analysis (based on ADF and Phillips-Perron tests) does not reject the null hypothesis of a unit root in each of our three series at the 10% level and points to nonstationarity, or an $I(1)$ structure for the series. Consequently, and following earlier studies, we proceed to a cointegration analysis of the pass-through. Although a linear analysis using the Johansen methodology provides some evidence of cointegration,[7] particularly between LIBOR and the mortgage rate, such an analysis of ignores the possibility of nonlinearity. The next section, therefore, discusses the econometric methodology we develop for dealing with possible nonlinear adjustment for interest rate changes, before results are considered in Section 5.

4. Econometric methodology

In common with many other studies, our pass-through analysis employs single equation modelling under the assumption of weak exogeneity. More specifically, we assume the official rate determined by the Bank of England is weakly exogenous to the market rate, which in turn is weakly exogenous to the mortgage rate, since banks' retail rates are not expected to affect market rate movements (de Bondt et al., 2005).

This section first discusses the ECM models on which our empirical analysis is based (subsection 4.1), with subsection 4.2 then outlining model specification and estimation. Our bootstrap testing methodology, which explicitly allows for the characteristics of the interest rate data, is detailed in subsection 4.3. A final subsection outlines the nature of the generalized impulse response functions later used to aid the interpretation of our estimated models.

4.1. Error-correction models

Assuming that all interest rates are (nonlinearly) cointegrated $I(1)$ variables, and with the exogeneity assumptions already noted, the threshold ECM (Balke and Fomby, 1997; Enders and Siklos, 2001;

[4] According to the Bank of England database, this series is the average of four major clearing banks' base rates and from May 1997 is identical to the database series for the official rate, except for the specific days when interest rates change. Although the official rate series in the database prior to May 1997 is typically 0.12 percentage points lower than this rate, for this earlier period the series we use is identical to that reported in the Bank of England's *Quarterly Bulletin* in relation to monetary policy decisions taken by the Chancellor and the Governor of the Bank.

[5] The data series for the average SVR of banks was discontinued after December 2007. To extend the series to the entire sample period, we regressed the average bank SVR on a constant and the available series for the combined mortgage rates of banks and building societies over January 1995 to December 2007. This estimated relationship was then used to generate fitted values for the banks' average SVR for January to August 2008. A check of this methodology applied to a shortened sample (and checked against actual SVR) confirms that it delivers satisfactory results. These results are available on request.

[6] If the degree and/or speed of the pass-through is not complete, an increase (decrease) in LIBOR will result in a decrease (increase) in MSPREAD because of small and/or slow responses of the mortgage rate.

[7] Detailed results for linear ECMs, including estimated models, can be found in Becker et al. (2010), which is a discussion paper version of the present paper.

Hansen and Seo, 2002) for the two stages of the pass-through has the form

$$\Delta libor_t = \sum_{i=1}^{p} \varphi_{1i} \Delta libor_{t-i} + \sum_{i=0}^{q} \vartheta_{1i} \Delta brate_{t-i} + \gamma_{11} M_{1t} u_{1t-1} + \gamma_{12}(1-M_{1t}) u_{1t-1} + v_{1t} \tag{1}$$

$$\Delta m\,rate_t = \sum_{i=1}^{p} \varphi_{2i} \Delta m\,rate_{t-i} + \sum_{i=0}^{q} \vartheta_{2i} \Delta libor_{t-i} + \gamma_{21} M_{2t} u_{2t-1} + \gamma_{22}(1-M_{2t}) u_{2t-1} + v_{2t} \tag{2}$$

where $u_{1t} = libor_t - \alpha_1 - \beta_1 brate_t$ and $u_{2t} = mrate_t - \alpha_2 - \beta_2 libor_t$ are the disequilibria at t in each of the two pass-through stages, M_{it} $(i=1,2)$ is the regime operating at time t for the i^{th} stage, and v_{1t}, v_{2t} are *iid* error terms with zero mean and constant variances. In an obvious notation, *brate*, *libor* and *mrate* indicate the base rate, LIBOR and the mortgage rate, respectively. The coefficients α_i and β_i $(i=1, 2)$ measure the mark-up (or down) and the degree of the pass-through in the long run, with complete pass-through indicated by $\beta_i = 1$ and incomplete pass-through by $\beta_i < 1$. The specification of (1) and (2) assumes that all effects of the base rate on the mortgage rate operate through the money market rate, an issue to which we return below.

The regimes for the nonlinear threshold ECM are specified through indicator variables, expressed as the Heaviside functions, such that

$$M_{it} = \begin{cases} 1 & z_{it} > \tau_i \\ 0 & z_{it} \le \tau_i \end{cases} \qquad i = 1,\ 2 \tag{3}$$

Even if the threshold variable z_{it} is known, the threshold value τ_i is typically unknown. This implies that nonstandard procedures are required for testing the presence of nonlinear cointegration between the interest rate pairs, which is the subject of the next subsection.

The threshold cointegration literature commonly adopts either the lagged disequilibrium or the changes in this disequilibrium as the threshold variable, corresponding to $z_{it} = u_{i,t-1}$ or $z_{it} = \Delta u_{i,t-1}$ for our case. The latter is referred to as M-TAR (momentum threshold autoregressive) adjustment by Enders and Siklos (2001), who suggest that it is appropriate when policy-makers smooth out large adjustments, and Payne (2007) adopts this specification when modelling the pass-through to retail interest rates in the United States. Sander and Kleimeier (2004), on the other hand, consider both possibilities, together with a band-TAR model represented by $z_{it} = |u_{i,t-1}|$, which implies that the speed of adjustment depends on the size of the disequilibrium. However, Hofmann and Mizen (2004) find that (actual or expected) changes in the official rate influence the speed of adjustment of retail to official rates in the UK, with two regimes dependent on $\Delta brate_t$ ($\Delta brate_t > 0$ versus $\Delta brate_t \le 0$). Nevertheless, this not only assumes a known zero threshold, but also conflates zero and negative base rate changes.

Due to differing views for possible nonlinear drivers in the interest rate pass-through, we examine each of $u_{i,t-1}$, $\Delta u_{i,t-1}$ and $|u_{i,t-1}|$. Further, a disequilibrium value $u_{1,t-1}$ could arise because either (or both) the base rate or LIBOR changes, suggesting that the underlying driver for adjustment may be the change that gives rise to the disequilibrium, namely $\Delta brate_{t-1}$ or $\Delta libor_{t-1}$. Hence we also consider each of these as the possible first-stage nonlinear drivers.[8] For analogous reasons, $\Delta brate_{t-1}$, $\Delta libor_{t-1}$ and $\Delta mrate_{t-1}$ are examined for the second stage of the pass-through. In addition, size effects are examined by considering the absolute values of these variables. This may be particularly important because the official rate frequently remains unchanged (as noted in Section 3), with a clear signal about monetary policy arguably being provided only in months when a rate change occurs. Similarly, small changes in LIBOR or the mortgage rate may be essentially noise and hence generate different adjustment responses compared to large changes.

[8] Contemporaneous $\Delta brate_t$ and its absolute value were also considered as the possible nonlinear driver for both (3) and (4), together with $\Delta libor_t$ for (4). However, stronger evidence of nonlinearity and cointegration were obtained using the lagged values of these variables.

4.2. Model specification and estimation

The lag orders in (1) and (2) are specified in the linear framework, excluding the equilibrium error terms. In particular, the Schwarz Bayesian criterion (SBC) is used in order to determine (separately) the lag orders p and q, up to a maximum lag order of 12 in each case and allowing intermediate lags to be dropped. These lags are then carried over to the threshold ECMs.

One issue in empirical modelling is the handling of "outlier" observations, which can play an important role in a nonlinear context. In this sense, a dummy variable is included to account for a residual whose (absolute) value is larger in magnitude than 3 standard errors. To maintain the asymptotic distribution of test statistics, relevant step dummies are added to the cointegrating relationship; see Doornik et al. (1998) for details. To ensure comparability of linear and nonlinear models, the same dummy variables are included in all models for a specific (first or second) stage of the pass-through, with these dummies observed from the linear model.

Stock (1987) recommends estimating linear ECMs using nonlinear least squares (NLS). This can be achieved by estimating initial values of the long-run coefficients using ordinary least squares (OLS), with the initial values of the parameters of the short-term dynamics then obtained by OLS conditional on these, with NLS finally applied to the whole model. Threshold ECMs are estimated by modifying the sequential least squares approach of Hansen (1997). That is, for each potential threshold value τ_i, which is typically in the middle 70% of the ordered values of the threshold variable, a threshold ECM is estimated through NLS conditional on this value, using the procedure just outlined for a linear ECM. The estimate $\hat{\tau}_i$ is the value minimizing the sum of squared residuals over these estimations. Estimates of the cointegrating vector and the remaining parameters are the NLS estimates associated with this $\hat{\tau}_i$.[9]

4.3. Testing for threshold cointegration

Prior to estimation of a threshold model, the presence of nonlinearity should be established. Although Balke and Fomby (1997) and Hansen and Seo (2002) undertake a test based on an initial linear cointegration analysis, Enders and Siklos (2001) argue that this is unsatisfactory due to the misspecification and low power of these tests in the presence of asymmetric adjustment. Instead, they propose a cointegration test that allows for threshold adjustment and, if cointegration is established, then test the null of symmetric adjustment using a standard F-test.

We follow Enders and Siklos (2001) by testing for the presence of cointegration allowing for asymmetric adjustment through the model

$$\Delta u_{it} = \gamma_{i1} M_{it} u_{it-1} + \gamma_{i2}(1-M_{it}) u_{it-1} + \sum_{j=1}^{q'} \delta_{ij} \Delta u_{it-j} + \eta_{it} \tag{4}$$

where u_{it} (i = 1,2) are as defined for (1) and (2), q' is the required number of lagged changes that ensures an *iid* structure for the error term, η_{it}, and the regimes for M_{it} are defined in (3). The null of no

[9] Although the disturbances (ν_{1t}, ν_{2t}) in (1)/(2) may be correlated, since each represents a "seemingly unrelated" system of equations, this is not taken into account in estimation (or the subsequent impulse response calculation), due to the complexity of the nonlinear procedure that is our principal focus. It may, however, be noted that we found the application of nonlinear ECMs to reduce this correlation substantially compared to a linear model.

cointegration, $\gamma_{i1}=\gamma_{i2}=0$, is tested against the alternative of threshold cointegration. As the threshold value, τ_i, defining M_{it} is unidentified under the null hypothesis, the test statistic sup LM_T^{nc} is obtained by maximization over the range of possible τ_i, defined as the central 70% of the distribution of the relevant z_{it}. The distribution of this test statistic is nonstandard and must be obtained by simulation.

Although Enders and Siklos (2001) provide critical values, these do not consider the possibility of a variable being discrete. Therefore, we develop a fixed design model-based bootstrap procedure along the lines of that suggested by Hansen and Seo (2002) in order to mimic the observed data features. Specifically, the bootstrap p-values for testing the null of no cointegration between LIBOR and the base rate are simulated through the following algorithm:

i) Estimate the long-run relationship $libor_t=\alpha_1+\beta_1 brate_t+u_{1t}$ by OLS; obtain $\hat{\alpha}_1$, $\hat{\beta}_1$.
ii) Generate the bootstrap DGP series $libor_t^*$ as

$$libor_t^*=\hat{\alpha}_1+\hat{\beta}_1 brate_t+\nu_{1t}^*, t=1,\ 2,\dots,T$$

where ν_{1t}^* is a random walk sequence with standard deviation set equal to the empirical residual standard deviation of u_{1t} and T is the sample size.
iii) Re-estimate the long-run relationship using $libor_t^*$ in conjunction with the actual $brate_t$ and obtain the residuals u_{1t}^*.
iv) Using the sequence u_{1t}^*, estimate the threshold model of (3) and (4), and calculate the bootstrap LM test statistic, $LM_T^{nc^*}(\tau_1)$, for the null of $\gamma_{11}=\gamma_{12}=0$ for each τ_1 on the grid set $[\tau_{1L}, \tau_{1U}]$, where τ_{1L} and τ_{1U} are the 15th and 85th percentiles of the potential threshold variable z_{1t}^*.
v) Obtain sup $LM_T^{nc^*}$ as

$$\sup LM_T^{nc^*}=\sup_{\tau_1\in[\tau_{1L},\tau_{1U}]} LM_T^{nc^*}(\tau_1).$$

vi) By repeating steps ii) to v), generate 50,000 bootstrap replications of sup $LM_T^{nc^*}$, and calculate the bootstrap p-value as the percentage of sup $LM_T^{nc^*}$ values that exceed the observed test statistic sup LM_T^{nc}.

It is straightforward to adapt this algorithm for the cointegration analysis between the mortgage rate and LIBOR, with $libor_t$ treated as exogenous. The only case where this procedure is not employed for cointegration testing is when the potential threshold variable is the absolute change in the base rate. In this case, given the infrequency with which base rate changes of more than 25 basis points are observed in our sample period, the only feasible threshold to be examined in (4) is zero, which is therefore known and no unobserved parameter problem arises.

When two interest rates are found to be cointegrated, the null hypothesis of symmetric adjustment, namely $\gamma_{i1}=\gamma_{i2}=\gamma_i$, is tested using a sup LM test. Although Enders and Siklos (2001) employ a standard F-test, conditional on the estimate of τ_i obtained from the cointegration testing, they note this could be problematic. In contrast, our approach continues to recognise that τ_i is unidentified under the null hypothesis and employs a model-based bootstrap procedure similar to Balke and Fomby (1997). However, since the sup LM test has a two-sided alternative, rejection of the null does not guarantee the stationarity of u_{it} in (4). As shown by Petrucelli and Woolford (1984) and Chan et al. (1985), necessary and sufficient conditions for stationarity are $\gamma_{i1}<0, \gamma_{i2}<0$ and $(1+\gamma_{i1})(1+\gamma_{i2})<1$.[10] Therefore, our procedure checks (for every bootstrap replication) that the estimated coefficients satisfy these stationarity conditions before testing the symmetry null hypothesis $\gamma_{i1}=\gamma_{i2}=\gamma_i$, and replications that do not satisfy stationarity are discarded.[11]

This bootstrap test procedure is used for all cases except when the base rate change is considered as the threshold variable, with asymptotic test statistics being employed in this case due to the known threshold of zero.

4.4. Dynamic analysis of threshold error-correction models

In order to provide further insights into the implications of the estimated nonlinear threshold cointegration models, generalized impulse response analysis is performed in relation to each of the two stages encapsulated in (1) and (2), and also for the system consisting of both equations.

Gallant et al. (1993) and Koop et al. (1996) point out that, unlike linear models, the impulse response function of a nonlinear model is not (in general) independent of either the history of the series at the time of the shock or the sign and size of the shock. Further, due to the analytical intractability of these models, the impulse response functions have to be obtained by simulation. In the interest rate pass-through literature, the only study utilizing impulse response analysis of a threshold ECM is Sander and Kleimeier (2004), who do not, however, take account of the history dependent nature of the impulse response functions.

In this study, we follow Koop et al. (1996) and define the generalized impulse response functions for the two-regime threshold ECMs in (1) and (2) as

$$\begin{aligned} &GI_Y(h,\nu_t,W_{t-1},X_{t+h}) \\ &\quad=E(Y_{t+h}|\nu_t,W_{t-1},X_{t+h})-E(Y_{t+h}|W_{t-1},X_{t+h}),\ \ h=0,\ 1,\dots,H \end{aligned} \tag{5}$$

where GI_Y is the generalised impulse response function of the variable Y, which is *libor* or *mrate* depending on the stage of pass-through under analysis, ν_t is an arbitrary shock applied at time t, W_{t-1} is the history (information set of all variables up to time $t-1$), X_{t+h} is the information set of weakly exogenous variables to $t+h$ and H is the horizon.[12]

More specifically, our threshold ECM models have two regimes, corresponding to $M_{it}=1$ and $M_{it}=0$, in (3). To examine the nature of the regime-dependent adjustment in (1) and/or (2), we compare the generalized impulse response functions for shocks occurring in each regime. For the interest rate pass-through to the money market, consider a set of k_1 occasions for which $M_{1t}=1$ and define W_{t-1} to be the corresponding set of k_1 sequences of initial (lagged) values of *libor* and *brate* required in (1), namely $libor_{t-p-1}$, …, $libor_{t-1}$, $brate_{t-q-1}$, …, $brate_{t-1}$. Similarly, for these same k_1 specific periods for which $M_{it}=1$, X_{t+h} is the corresponding set of k_1 sequences of values $brate_t$, …, $brate_{t+h}$.

To calculate the generalized impulse response function in (5) conditional on $M_{it}=1$, we simulate Y forward from all k_1 histories, by randomly drawing innovations from the empirical distributions of estimated model residuals. The difference between a particular simulation $Y_{t+h}|\nu_t$, W_{t-1}, X_{t+h} and $Y_{t+h}|W_{t-1}$, X_{t+h} is the additional (given) perturbation ν_t. The generalised impulse response function (conditional on $M_{it}=1$) is then obtained by first averaging across 10,000 simulations for every particular history and subsequently averaging across all k_1 histories for which $M_{1t}=1$.

Generalized impulse response functions for the regime corresponding to $M_{it}=0$ and for the regimes in (2) are obtained in an analogous way. Impulse response functions are also presented when the

[10] If the transition variable z_{it} is chosen to be an absolute value we only require $\gamma_{i1}<0$ for global stationarity.

[11] The steps of our procedure are detailed by Becker et al. (2010).

[12] Both Gallant et al. (1993) and Koop et al. (1996) examine impulse response functions of nonlinear autoregressive models. We modify their approach for nonlinear univariate ECMs by assuming that the weakly exogenous variables are known to time $t+h$.

Table 1
Threshold cointegration test results.

Nonlinear Driver	Pass-through to LIBOR ($i=1$) To January 2006 Cointegration	Asymmetry	To August 2008 Cointegration	Asymmetry	Pass-through to mortgage rate ($i=2$) To January 2006 Cointegration	Asymmetry	To August 2008 Cointegration	Asymmetry
$\hat{u}_{i,t-1}$	13.623	4.263	11.493	4.919	14.511	2.549	20.079	2.967
	[0.002]**	[0.128]	[0.069]*	[0.092]*	[0.013]**	[0.343]	[0.001]**	[0.268]
$\Delta\hat{u}_{i,t-1}$	16.427	7.346	11.024	NA	17.977	5.419	22.466	4.646
	[0.018]**	[0.035]**	[0.172]		[0.008]**	[0.118]	[0.001]**	[0.189]
$\Delta libor_{t-1}$	27.238	19.231	16.313	10.095	19.998	7.665	21.598	3.667
	[0.000]**	[0.000]**	[0.021]**	[0.009]**	[0.004]**	[0.033]**	[0.001]**	[0.298]
$\Delta brate_{t-1}$	12.017	2.498	15.078	9.733	18.133	5.593	17.923	6.649
	[0.032]**	[0.138]	[0.004]**	[0.052]*	[0.002]**	[0.019]**	[0.000]**	[0.011]**
$\Delta mrate_{t-1}$					18.019	5.465	22.365	4.532
					[0.006]**	[0.073]*	[0.001]**	[0.163]
$\lvert\hat{u}_{i,t-1}\rvert$	18.990	10.164	17.811	11.704	15.202	3.311	20.078	2.965
	[0.006]**	[0.028]**	[0.018]**	[0.017]**	[0.022]**	[0.594]	[0.003]**	[0.591]
$\lvert\Delta\hat{u}_{i,t-1}\rvert$	25.173	16.961	11.860	NA	15.934	3.148	21.421	3.466
	[0.000]**	[0.000]**	[0.120]		[0.016]**	[0.373]	[0.001]**	[0.373]
$\lvert\Delta libor_{t-1}\rvert$	22.476	13.995	17.919	11.819	23.039	11.043	25.605	8.185
	[0.001]**	[0.002]**	[0.004]**	[0.005]**	[0.000]**	[0.006]**	[0.000]**	[0.028]**
$\lvert\Delta brate_{t-1}\rvert$	9.793	0.053	8.468	1.672	25.169	13.410	30.806	14.051
	[0.007]**	[0.818]	[0.014]**	[0.195]	[0.000]**	[0.000]**	[0.000]**	[0.000]**
$\lvert\Delta mrate_{t-1}\rvert$					20.074	7.750	24.577	7.026
					[0.004]**	[0.022]**	[0.001]**	[0.045]**

Notes: The tests for (possibly nonlinear) cointegration and asymmetry are described in subsection 4.3. The required number of lagged changes to ensure *iid* residuals in (4) is two for all cases where the pass-through is to LIBOR ($i=1$) and zero for all cases of the pass-through to the mortgage rate ($i=2$). The values in brackets in the table are p-values. Except when $|\Delta brate_{t-1}|$ is considered as the threshold variable, the p-values are obtained using the bootstrap algorithms (with 50,000 replications) described in subsection 4.3. For the threshold variable $|\Delta brate_{t-1}|$, the threshold value is set to 0 and chi-square p-values are reported. NA indicates that asymmetry test is not reported due to lack of evidence for cointegration.

two stages of the pass-through are considered (base rate shocks being transmitted to the mortgage market via the money market). In this case, four regimes are possible for (M_{1t}, M_{2t}), since different regimes can apply for each of the stages.

5. Estimation results

After discussing cointegration test results (subsection 5.1), the estimated threshold ECM models are presented in the following two subsections.[13] While no diagnostic test results are shown in order to conserve space, all estimated ECM models easily pass conventional tests for residual autocorrelation and conditional heteroscedasticity.[14] The implications of the estimated models, including the generalized impulse response functions, are considered in Section 6.

5.1. Cointegration tests

Using the testing methodology detailed in Section 4.3, together with the potential nonlinear drivers discussed in Section 4.1, Table 1 presents the threshold cointegration results for both steps of the pass-through, estimated over both samples (to January 2006 and August 2008).

Results in the left-hand panel of Table 1, for the pass-through from base rates to LIBOR, provide evidence for cointegration over the main sample period irrespective of the potential nonlinear driver examined. However, the strength of this evidence (in terms of p-values) differs. Nevertheless all drivers based on magnitude, either of the disequilibrium or of the change in LIBOR or the base rate, reject the null hypothesis of no cointegration null with p-values of less than 1%, with similar result for $\Delta libor_{t-1}$. Evidence of nonlinear cointegration continues to apply when the longer period is considered, with this again significant at 1% for the threshold variable $|\Delta libor_{t-1}|$. Further, for both sample periods, the associated test for asymmetric adjustment is highly significant. Taken overall, this evidence points to conditions in the money market itself being crucial for the nature of adjustment towards the long-run equilibrium of LIBOR with the base rate, with different rates of adjustment to equilibrium for larger versus smaller changes in the money market rate.

For the pass-through from LIBOR to the mortgage rate (right-hand panel of Table 1), cointegration is confirmed irrespective of the potential threshold variable considered, often with very small p-values. For the sample to January 2006, the strongest evidence of cointegration applies when the magnitude of changes in LIBOR or the base rate is considered as the nonlinear driver, with these also providing strong evidence of nonlinear adjustment. This is largely confirmed by the longer sample, although that points more clearly to the base rate variable as nonlinear driver for this stage of adjustment. Thus, it appears that retail mortgage rates adjust to their equilibrium with LIBOR at different rates, depending on the magnitude of change in the Bank of England's base rate.

One general implication of the results in Table 1 is that both stages of the interest rate pass-through depend on whether interest rates are relatively stable or not. This indicates that previous literature which follows the Enders and Siklos (2001) framework in assuming the nonlinear driver to be a cointegration residual or its change (or, as in Hofmann and Mizen, 2004, changes in the base rate), apparently overlooks the potentially most important source of nonlinearity, namely the magnitude of rate changes. This finding is in line with the argument of Balke and Fomby (1997) that the costs of adjustment can lead to asymmetric adjustment to equilibrium, with the nature of this asymmetry for UK rates examined in the following subsections.

5.2. Pass-through to LIBOR

Table 2 provides estimation results for the pass-through from base rates to LIBOR, using a nonlinear ECM with threshold variable $|\Delta libor_{t-1}|$. Although a number of possible nonlinear drivers lead to evidence of cointegration in Table 1, $|\Delta libor_{t-1}|$ is selected as it yields the best fit (according to SBC, AIC and the residual standard error) for

[13] Results for linear models, in addition to the threshold ECMs, can be found in the discussion paper version of this paper (Becker et al., 2010).

[14] All models considered (for both sample periods) pass these tests at the 10% level, except the threshold ECM for the pass-through to the mortgage rate (for the extended sample to 2008) yields an ARCH test p-value of 0.099.

Table 2
Estimated threshold models for pass-through to LIBOR.

	Estimated to			Estimated to	
	January 2006	August 2008		January 2006	August 2008
Cointegrating relation (dependent variable $libor_t$)			Short-run adjustment (dependent variable $\Delta libor_t$)		
Constant	0.139	0.424	$M_{1t} \times \hat{u}_{1,t-1}$	−0.513	−0.442
	[0.496]	[0.021]		[0.001]	[0.001]
$brate_t$	0.990	0.948	$(1 - M_{1t}) \times \hat{u}_{1,t-1}$	0.044	0.056
	[0.000]	[0.000]		[0.392]	[0.281]
$D0001$	−0.080	−0.087	$\Delta brate_t$	0.847	0.844
	[0.150]	[0.133]		[0.000]	[0.000]
$D0110$	0.132		$\Delta brate_{t-1}$	0.174	0.369
	[0.116]			[0.000]	[0.000]
$D0708$		0.120	$\Delta brate_{t-2}$		0.216
		[0.284]			[0.000]
$D0801$		0.025	$\Delta brate_t \times d0708_t$		−0.269
		[0.867]			[0.238]
Complete pass-through	0.092	3.395	$\Delta brate_{t-1} \times d0708_t$		−0.708
	[0.763]	[0.065]			[0.001]
			$\Delta brate_{t-2} \times d0708_t$		−0.252
					[0.199]
Model statistics					
$\hat{\sigma}$	0.069	0.067	$\Delta libor_{t-1}$		−0.182
					[0.000]
			$\Delta libor_{t-2}$		−0.162
					[0.005]
			τ	0.090	0.090

Notes: *Dyymm* indicates a step dummy for month mm of year yy; the short-run adjustment equation includes the corresponding impulse dummy variables, *dyymm* (coefficients not shown). The threshold models use $|\Delta libor_{t-1}|$ as the nonlinear driver (see text). All values in brackets are *p*-values; for coefficients these test the null hypothesis of zero while the complete pass-through test is a Wald test of the null hypothesis that the coefficient on $brate_t$ in the long-run model is unity.

Table 3
Estimated models for pass-through to mortgage rate.

	Estimated to			Estimated to	
	January 2006	August 2008		January 2006	August 2008
Cointegrating relation (dependent variable $mrate_t$)			Short-run adjustment (dependent variable $\Delta mrate_t$)		
Constant	2.227	2.239	$M_{2t} \times \hat{u}_{2,t-1}$	−0.702	−0.751
	[0.000]	[0.000]		[0.000]	[0.000]
$libor_t$	0.829	0.828	$(1 - M_{2t}) \times \hat{u}_{2,t-1}$	−0.073	−0.072
	[0.000]	[0.000]		[0.000]	[0.005]
$D9706$	0.376	0.364	$\Delta libor_t$	0.025	0.019
	[0.005]	[0.000]		[0.498]	[0.580]
$D9812$	−0.102	−0.093	$\Delta libor_{t-1}$	−0.006	−0.020
	[0.192]	[0.187]		[0.919]	[0.713]
$D9903$	0.115	0.119	$\Delta libor_{t-2}$	0.160	0.148
	[0.128]	[0.083]		[0.000]	[0.000]
$D0612$		0.134	$\Delta libor_t \times d0612_t$		0.179
		[0.001]			[0.008]
$D0805$		−0.074	$\Delta libor_{t-1} \times d0612_t$		0.096
		[0.865]			[0.209]
Complete pass-through	54.343	67.388	$\Delta libor_{t-2} \times d0612_t$		−0.036
	[0.000]	[0.000]			[0.583]
Model statistics					
$\hat{\sigma}$	0.054	0.053	τ	0	0

Notes: *Dyymm* indicates a step dummy for month mm of year yy; the short-run adjustment equation includes the corresponding impulse dummy variables, *dyymm* (coefficients not shown). Both threshold models use $|\Delta brate_{t-1}|$ as the nonlinear driver (see text). All values in brackets are *p*-values; for coefficients these test the null hypothesis of zero while the complete pass-through test is a Wald test of the null hypothesis that the coefficient on $libor_t$ in the long-run model is unity.

models estimated over both periods.[15] The left-hand panel shows estimates of the long-run equilibrium relationship, while parameters relating to the short-run dynamics are in the right-hand panel. To conserve space, the short-run coefficients associated with impulse dummy variables are not shown, but these are always individually significant at levels of significance of 5% or (typically) less.

The ECM estimated over the reference sample (to January 2006) is compatible with the pass-through to LIBOR being complete in the long run. Indeed, the estimated coefficient of $brate_t$ is very close to unity and the respective hypothesis test has a large *p*-value of 0.763.[16] Noting that values in square brackets are *p*-values for a null hypothesis of zero, the mark-up (captured by the intercept) is not significantly different from zero, supporting the graphical evidence in Fig. 1.

The short-run dynamics indicate that much of the pass-through is immediate and implies that when LIBOR changes by more than around ± 0.1 percentage points in any month, there is a further adjustment in the next period to remove half of the resulting disequilibrium. As a careful inspection of Fig. 1 makes clear, LIBOR sometimes anticipates base rate changes, which provides a rationale for why the occurrence of nontrivial changes in the money market rate is the driver for the nonlinear ECM specification. On the other hand, when LIBOR changes are very small, the adjustment coefficient is not significantly different from zero, with the small changes in LIBOR presumably reflecting very short run and minor fluctuations in the money market.

Extending the sample to 2008 gives rise to a number of changes. There is now less evidence of complete pass-through and the two dummy variables, although not individually significant here, seem to drive a lasting wedge between the base and LIBOR rates (consistent with what can be gleaned from Fig. 1); this is also indicated by the significant mark-up. The August 2007 hike in the LIBOR rate, which was not mirrored by any increase in the base rate, was the start of a period in which the LIBOR rate persistently exceeded the base rate, and this is also reflected in different dynamic responses to base rate changes from that date.[17] Nevertheless, there is little change in the disequilibrium adjustment.

It is interesting to examine events underlying the dummy variables identified in these specifications. Of these, the January 2000 dummy ($D0001$) corresponds to millennium effects which (although details are not shown) are highly significant in all short-run specifications. The $D0708$ dummy (August 2007) corresponds to the beginning of the Northern Rock crisis, which resulted in the nationalization of Northern Rock in January 2008 ($D0801$).

5.3. Pass-through from LIBOR to mortgage rate

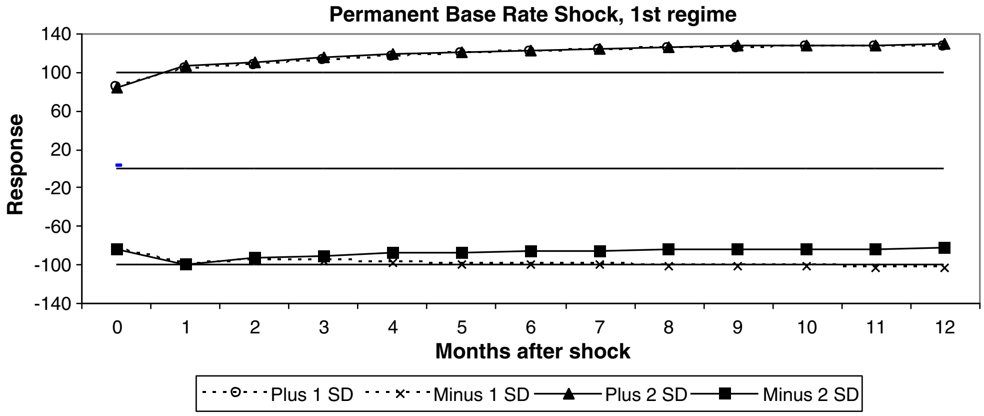
Results for the estimated ECM models for the pass-through from LIBOR to mortgage rates are shown in Table 3. These employ $|\Delta brate_{t-1}|$ as the nonlinear driver, since this leads to the strongest evidence of threshold cointegration and asymmetry in Table 1 over both periods.[18] In the light of the infrequency of base rate changes of more than 25 basis points, the only feasible threshold value is zero (predetermined),

[15] Specifically, $|\Delta\hat{u}_{1t-1}|$, $\Delta libor_{t-1}$ and $|\Delta libor_{t-1}|$ were investigated over the shorter sample, since these yield similar values, with very high significance, for both the cointegration and asymmetry tests. For the same reason, $\Delta libor_{t-1}$ and $|\Delta libor_{t-1}|$ were considered over the extended sample.

[16] None of the step dummy variables included in the longrun specification are significant (at 5%), which is compatible with this capturing enduring features of the relationship between money market and base rates. They remain in the specification as their impulse dummy counterparts in the shortrun dynamics are significant.

[17] Tests of stability did not indicate a change in the coefficients for $\Delta libor_{t-i}$ $(i = 1, 2)$ at this date.

[18] For the reference sample, an ECM model with $|\Delta libor_{t-1}|$ as the threshold variable was also estimated, but that using $|\Delta brate_{t-1}|$ yielded the best fit according to SBC, AIC and the residual standard error.



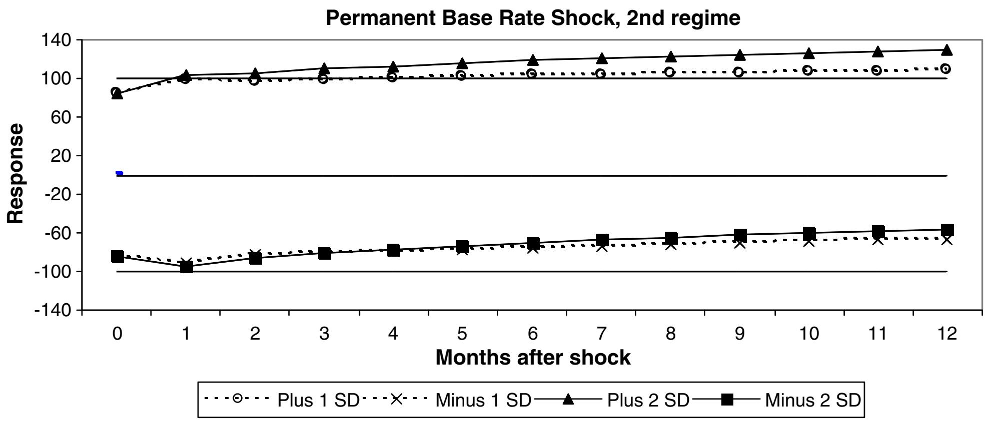


Fig. 3. Dynamic responses of LIBOR to base rate shocks. Responses are given in percentage terms and are obtained through stochastic simulations of the two-regime threshold ECM using 10,000 replications. 1st and 2nd regimes refer to cases where $|\Delta libor_{t-1}|>0.090$ and $|\Delta libor_{t-1}|\leq 0.090$, respectively.

and hence the regimes separate months where the rate remains constant versus those where the base rate changes.

Notice, first, that complete pass-through is clearly rejected for both sample periods, with $libor_t$ having a long-run coefficient of 0.83. Hence, although base rate changes are fully passed through to the money market rate in the long-run, they are incompletely passed through to the mortgage rate. Also, but not surprisingly, both models provide evidence of a significant mark-up of the retail mortgage rate over LIBOR.

The responses to disequilibrium are also interesting. When a base rate change occurred in the preceding month ($M_{2t}=1$), there is fast disequilibrium correction, with estimated adjustment coefficients of -0.70 and -0.75 for the two samples. On the other hand, stability in the monetary policy stance and no change in base rates ($M_{2t}=0$) is associated with very sluggish, yet statistically significant, adjustment (coefficient -0.07). Thus, disequilibria arising from LIBOR movements that are backed by changes in the monetary policy instrument are eliminated relatively quickly, whereas those arising in a stable monetary environment are not.[19] Although a change in dynamics is indicated at the end of 2006, and impulse dummy variables are required to account for specific events during a period of severely stressed market conditions, it is remarkable that the essential results carry over to a sample period ending in August 2008.

Although it is difficult to pin down the events associated with the December 2006 dummy (and changed dynamics) to a particular event, it roughly coincides with the market's realisation that the decline in house price inflation, which began in the summer 2006, would be long lasting. The dummy for June 1997 ($D9706$) coincides with the conversion of Alliance & Leicester and Halifax from building societies to banks, which may have led to decreased competition between building societies and banks,[20] resulting in an increase in the mark-up of banks' mortgage rates.

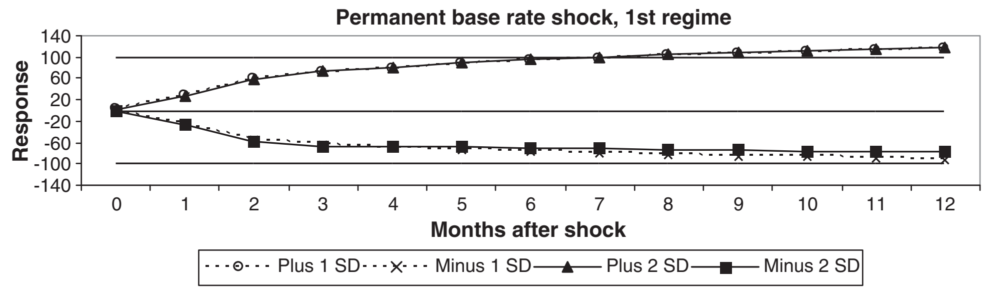
6. Impulse response analysis and discussion

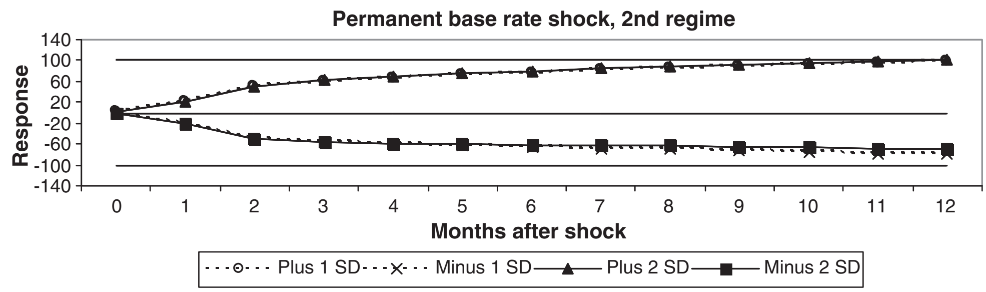
In this section we will present generalized impulse response functions (GIRF, see Section 4.4) for our pass-through models and use these to comment on the nature of the two-stage mechanism. GIRFs are computed for a horizon (h) of 12 months for shocks of $\pm 1\hat{\sigma}$ and $\pm 2\hat{\sigma}$, where $\hat{\sigma}$ is the relevant estimated residual standard deviation. These are regime-dependent, with the regime being that applying at the period of the initial shock. All results employ the coefficient estimates of the reference sample period (to January 2006). A GIRF analysis undertaken for the extended sample ending in August 2008 shows very similar results and hence are not reported.

Fig. 3 examines the pass-through of a base rate shock to LIBOR, where the increase or decrease in the policy rate is permanent. Regimes defined by $|\Delta libor_{t-1}|>0.090$ are in the upper panel and $|\Delta libor_{t-1}|\leq 0.090$ in the lower panel; the vertical axis shows the percentage of the initial shock that is adjusted at a given horizon. In both regimes, negative shocks are generally not fully passed through, particularly when these occur in the second regime when the money market is characterized by small changes in the previous period. Hence base rate reductions do not fully translate to the money market rate when

[19] Although results are not shown, it is interesting to note that allowing for the nonlinear equilibrium correction substantially improves the fit compared to a linear specification, with the residual standard error being reduced by more than a quarter; see Becker et al. (2010).

[20] See Heffernan (2005) for evidence suggesting that post conversion building societies adjusted their price setting behaviour to look more like that of a normal bank.





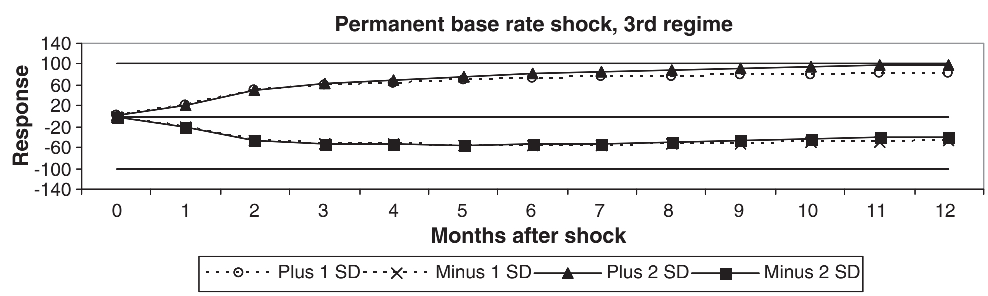


Fig. 4. Dynamic responses of the mortgage rate to base rate shocks. Responses are given in percentage terms and are obtained through stochastic simulations of the two-stage, two-regime threshold ECM using 10,000 replications. Regimes refer to cases where: 1st regime: $|\Delta libor_{t-1}|>0.090$ and $|\Delta brate_{t-1}|>0$; 2nd regime: $|\Delta libor_{t-1}|>0.090$ and $|\Delta brate_{t-1}|=0$; 3rd regime: $|\Delta libor_{t-1}|\leq 0.090$ and $|\Delta brate_{t-1}|=0$. The regime where $|\Delta libor_{t-1}|\leq 0.090$ and $|\Delta brate_{t-1}|>0$ is not considered due to an insufficient number of initial values.

they do occur in a previously stable market. On the other hand, positive shocks have a stronger than one-to-one effect on the LIBOR rate. Sander and Kleimeier (2004) and de Bondt (2005) explain overshooting to positive base rate shocks as indicating that banks increase their risk premium due to potentially increased default risk, which may apply to the money market rate in our case. This overshooting applies irrespective of the size of the base rate increase in Regime 1, but only to large increases in Regime 2.

GIRFs computed for mortgage rate responses to LIBOR shocks (and presented in Becker et al., 2010) show that the mortgage rate reacts equally to positive and negative LIBOR rate shocks, which occurs because regimes are governed by the base rate. In contrast to the first stage of the pass-through, the second-stage adjustment is sluggish especially in the regime characterized by an unchanged base rate in $t-1$ (after one year less than 80% of the change has been passed through).

The pass-through of monetary policy, as encapsulated in the base rate, to the mortgage rate faced by households is of particular interest, since it gives an insight into how monetary policy decisions affect consumers. In our two-stage model, the effects of base rate shocks are transmitted first to the LIBOR rate via the threshold cointegration model of Table 2, with the simulated LIBOR rates then used as the histories of money market rates relevant for the mortgage rate in the specification of Table 3.[21] As these two models have distinct threshold variables, four different regimes may be implied. Of these four, three regimes are analyzed in Fig. 4; first the policy change regime (Regime 1) in which period $t-1$ is characterized by a change in the base rate and a nontrivial change in the LIBOR rate (greater than 0.09 percentage points in magnitude). The second regime is a LIBOR only change regime, which represents the case in which the LIBOR rate changed nontrivially (i.e., $|\Delta libor_{t-1}|>0.090$), despite there being no change in the base rate. The third regime is the stable regime which is defined by no (nontrivial) change in either the base rate or LIBOR. The potential fourth regime (base rate change but no corresponding change in the LIBOR) yields only three empirical observations over the sample period, and hence is not considered to be empirically plausible.

Comparing responses across regimes in Fig. 4, it is clear that the mortgage rate adjusts more quickly when the base rate shock occurs after previous changes in the monetary policy stance (Regime 1) than otherwise. This is compatible with mortgage providers facing lower costs of adjustment in a context of recent previous base rate changes, because their systems are prepared for further change.

Asymmetries can also be observed in relation to the sign of the monetary policy shock. Positive base rate shocks are fully reflected in mortgage rates between 6 months (policy change Regime) and 10 months (Regimes 2 and 3) after the occurrence of the shock. Negative base rate shocks, however, fail to be fully transmitted to a corresponding decline in mortgage rates. These asymmetries are strongest in Regimes 2 and 3, which represent cases in which a monetary shock is not preceded by another change in the base rate. The lowest proportion of about one half of the base rate decline is ultimately passed on to mortgage holders in the stable regime (Regime 3). Hence the

[21] The correlation between the residuals of the threshold ECMs of Tables 2 (Stage 1) and 3 (Stage 2) is assumed to be zero; this is a reasonable approximation, given an empirical correlation of −0.230 between these residuals.

effectiveness of monetary policy that acts to stimulate consumer spending through declines in the base rate is substantially impeded in this regime due to the lower portion of the decline being passed to the mortgage rate. Comparing Regimes 1 and 3, these results imply that a series of smaller base rate declines (Regime 1) are more effective in this sense than a single larger base rate shock occurring after a stable period. Although the size effects are not very large, Fig. 4 (in common with Fig. 3) implies that a larger proportion of base rate declines are passed on when the declines are small (one compared with two standard deviation shocks).

These results shed important light on the previous finding of asymmetries in the mortgage market in Fuertes et al. (2010). There the asymmetries are attributed to the structure of the mortgage market. The analysis presented here, however, implies that important nonlinearities arise also in the first-step of our pass-through process and hence ought to be explained in the interbank rather than the mortgage market.

7. Conclusions

This paper investigates the transmission of interest rate shocks induced by monetary policy to the mortgage rate. In order to dissect interest shocks appropriately, the transition is separated into two steps (from the base rate to LIBOR and from LIBOR to mortgage rates), allowing for the possibility of asymmetries in both steps. This reveals that asymmetries which appear to be in the mortgage market (namely, incomplete pass-through of base rate reductions to the mortgage rate, but complete pass-through of base rate increases, over a one year horizon), are primarily a feature of the money market rather than the mortgage market itself.

Nonlinearities play an important role in our analysis, as adjustment speeds to long-run equilibria vary significantly depending on an underlying state variable. In general, however, we find that adjustment speeds are significantly greater when interest rate movements are motivated by clear monetary policy signals. The nonlinear analysis further reveals that the interest rate pass-through between the policy rate and the money market is complete, but that the pass-through from the money market to the mortgage market is not. An extended sample, reaching into the beginning of the recent credit crunch period, provides a robustness check on this analysis.

From the perspective of the effectiveness of monetary policy, our results imply that tightening policy through base rate increases will have a relatively quick impact on consumers through an increase in mortgage rates and a consequent reduction in their spending power. Since decreases are not fully passed on, base rate declines do not have the opposite effect of increasing consumers' spending power to a comparable extent. The latter statement is particularly relevant in the stable regime where no base rate change applied in the previous month and any change last month in LIBOR was very small. Since a greater proportion of a base rate decrease is passed to mortgage rates when change also occurred in the previous month, while the mortgage rate response to a base rate increase is also quicker, a policy of interest rate smoothing by the central bank facilitates the pass-through process. Consequently, it is anticipated that interest rate smoothing will enhance the effectiveness of monetary policy.

The modelling approach adopted in this paper, in addition to allowing for nonlinear cointegration between the different interest rates, also includes a novel approach to statistical inference by explicitly allowing for the discrete nature of base rate changes. Indeed, a general feature of our approach is the extensive use made of bootstrap inference, which is employed for testing the presence of both cointegration and nonlinearity.

References

Balke, N.S., Fomby, T.B., 1997. Threshold cointegration. International Economic Review 38, 627–645.

Becker, R., Osborn, D.R., Yildirim, D., 2010. A threshold cointegration analysis of interest rate pass-through to UK mortgage rates. Discussion paper number 141. Centre for Growth and Business Cycles Research, University of Manchester.

Benati, L., 2004. Evolving post-World War II UK economic performance. Journal of Money, Credit, and Banking 36, 691–717.

Burgstaller, J., 2005. Interest rate pass-through estimates from vector autoregressive models. Working Paper No. 0510. Johannes Kepler University of Linz, Austria.

Chan, K.S., Petrucelli, J., Tong, H., Woolford, S., 1985. A multiple-threshold AR(1) model. Journal of Applied Probability 22, 267–279.

Corvoisier, S., Gropp, R., 2001. Bank concentration and retail interest rates. ECB Working Paper No. 72. European Central Bank.

de Bondt, G., 2005. Interest rate pass-through: empirical results for the euro area. German Economic Review 6, 37–78.

de Bondt, G., Mojon, B., Valla, N., 2005. Term structure and the sluggishness of retail bank interest rates in euro area countries. ECB Working Paper No. 518. European Central Bank.

Doornik, J.A., Hendry, D.F., Nielsen, B., 1998. Inference in cointegrating models: UK M1 revisited. Journal of Economic Surveys 12, 533–572.

Enders, W., Siklos, P.L., 2001. Cointegration and threshold adjustment. Journal of Business and Economic Statistics 19, 166–176.

Fuertes, A., Heffernan, S.A., Kalotychou, E., 2010. How do UK banks react to changing central bank rates? Journal of Financial Services Research 37, 99–130.

Gallant, A.R., Rossi, P.E., Tauchen, G., 1993. Nonlinear dynamic structures. Econometrica 61, 871–907.

Hansen, B., 1997. Inference in TAR models. Studies in Nonlinear Dynamics and Econometrics 2, 1–14.

Hansen, B., Seo, B., 2002. Testing for two-regime threshold cointegration in vector error-correction model. Journal of Econometrics 110, 293–318.

Heffernan, S.A., 1993. Competition in British retail banking. Journal of Financial Services Research 7, 309–332.

Heffernan, S.A., 1997. Modelling British rate adjustment: an error correction approach. Economica 64, 211–231.

Heffernan, S.A., 2005. The effects of UK building society conversion on pricing behaviour. Journal of Banking and Finance 29, 779–797.

Hofmann, B., Mizen, P., 2004. Interest rate pass-through and monetary transmission: evidence from individual financial institutions' retail rates. Economica 71, 99–123.

Kesriyeli, M., Osborn, D.R., Sensier, M., 2006. Nonlinearity and structural change in interest rate reaction functions for the US, UK and Germany. In: Milas, C., Rothman, P., van Dijk, D. (Eds.), Nonlinear Time Series Analysis of Business Cycles. Elsevier, pp. 283–310.

Koop, G., Pesaran, M.H., Potter, S.M., 1996. Impulse response analysis in nonlinear multivariate models. Journal of Econometrics 74, 119–147.

Kuttner, K.N., 2001. Monetary policy surprises and interest rates: evidence from the Fed funds futures market. Journal of Monetary Economics 47, 523–544.

Miles, D., 2004. The UK Mortgage Market: Taking a Longer-Term View. Report to the UK Treasury. published at www.hm-treasury.gov.uk/consult_miles_index.htm.

Neumark, D., Sharpe, S., 1992. Market structure and the nature of price rigidity: evidence from the market for consumer deposits. Quarterly Journal of Economics 107, 657–680.

Paisley, J., 1994. A model of building society interest rate setting. Bank of England Working Paper No. 22. Bank of England.

Payne, J.E., 2006a. The response of the conventional mortgage rate to the federal funds rate: symmetric or asymmetric adjustment. Applied Financial Economics Letters 5, 279–284.

Payne, J.E., 2006b. More on the monetary transmission mechanism: mortgage rates and the federal funds rate. Journal of Post Keynesian Economics 29, 247–257.

Payne, J.E., 2007. Interest rate pass-through and asymmetries in adjustable rate mortgages. Applied Financial Economics 17, 1369–1376.

Petrucelli, J., Woolford, S., 1984. A threshold AR(1) model. Journal of Applied Probability 21, 270–286.

Sander, H., Kleimeier, S., 2004. Convergence in euro-zone retail banking? What interest rate pass-through tells us about monetary policy transmission, competition and integration. Journal of International Money and Finance 23, 461–492.

Sarno, L., Thornton, D.L., 2003. The dynamic relationship between the federal funds rate and the Treasury bill rate: an empirical investigation. Journal of Banking and Finance 27, 1079–1110.

Stephens, M., 2007. Mortgage market deregulations and its consequences. Housing Studies 22, 201–220.

Stock, J.H., 1987. Asymptotic properties of least squares estimators of cointegrating vectors. Econometrica 55, 1035–1056.